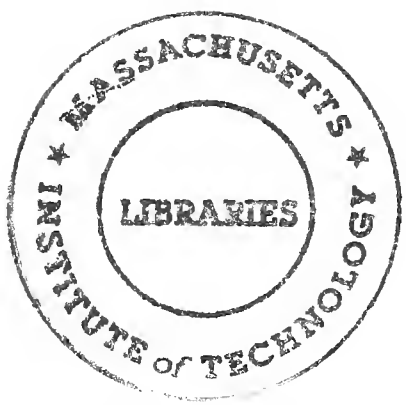
MASSACHUSETTS INSTITUTE of TECHNOLOGY
LIBRARIES

working paper
department
of economics

HEALTH INSURANCE AVAILABILITY AND THE RETIREMENT DECISION

Jonathan Gruber
Brigitte C. Madrian

94-4 Aug. 1993

massachusetts
institute of
technology

50 memorial drive
cambridge, mass. 02139

HEALTH INSURANCE AVAILABILITY AND THE RETIREMENT DECISION

Jonathan Gruber
Brigitte C. Madrian

94-4 Aug. 1993

Health Insurance Availability and the Retirement Decision

Jonathan Gruber
MIT and NBER

Brigitte C. Madrian
Harvard University and NBER

August, 1993

We wish to thank David Cutler, Henry Farber, Robin Lumsdaine, Jim Poterba, Andrew Samwick, Jim Stock, Richard Zeckhauser, and participants at seminars at Harvard, MIT, and the NBER for their comments, and Aaron Yelowitz for his help in assembling the SIPP data. We also acknowledge financial support from the James Phillips Fund, the Harry and Lynde Bradley Foundation, and the National Institute of Aging.

ABSTRACT

Because individuals aged 55-64 face large and uncertain medical expenditures without the guarantee of public insurance coverage provided by Medicare, the availability of post-retirement health insurance could be an important determinant in the retirement decisions of this group. We investigate the effect of health insurance on retirement by focusing on state and federal "continuation of coverage" mandates which grant the retiree the right to continue purchasing health insurance through a previous employer for a specified number of months after leaving the firm. We exploit variation in the timing and generosity of these laws to identify the effect of the availability of continuation coverage on retirement decisions, using data on 55-64 year-old males from the Current Population Survey and the Survey of Income and Program Participation. We find a sizeable and significant effect of continuation coverage on retirement; one year of mandated continuation benefits raises retirement rates by 20%. The effect appears to be uniform at all ages rather than larger near the age of Medicare eligibility. There is also a large increase in the insurance coverage of individuals who would have retired in the absence of continuation benefits. Our findings have important implications for policies which change the insurance coverage of early retirees, such as national health insurance.

The vast majority of Americans receive their health insurance coverage through one of two sources. Among the non-elderly, the primary source of coverage is their own or a family member's employment. For those over 65, Medicare covers a substantial fraction of medical expenditures. One group of individuals which often finds itself excluded from both of these systems are those who retire before the age of 65. Furthermore, medical expenditures for 55-64 year-olds are both high and uncertain, so that neither individual health insurance nor self-insurance of medical expenditures may be very attractive options for early retirees. Therefore, it seems likely that the availability of post-retirement health insurance coverage could be an important determinant of the retirement decision for older workers.

If the timing of retirement is very sensitive to insurance coverage, it could have major implications for the costs and consequences of a number of current policy proposals such as postponing the age of Medicare eligibility to 67 or providing government-sponsored health insurance to all individuals. Despite its likely importance, however, the role of health insurance coverage has, until recently, been largely ignored in the literature on retirement behavior. Given the sizeable evidence that health status is an important determinant of retirement, this is a surprising omission.

In this paper, we attempt to address this gap by examining the effect of state and federal "continuation of coverage" mandates on the retirement decision. These mandates grant individuals the right to continue purchasing health insurance through a previous employer for a specified number of months after leaving the firm. Although individuals must pay the average employer cost of their group insurance, the price to the early retiree is typically well below that of a policy purchased in the individual market. This is both because of high loading factors on individual insurance, and because individual policies are generally age-rated while the cost of

continuation coverage is not. By reducing the cost of health insurance after retirement, continuation of coverage benefits may increase the likelihood of early retirement for those individuals whose employers do not already provide post-retirement health insurance. The differences across states and over time in the availability and generosity of these laws provides the variation needed to assess the sensitivity of retirement behavior to this type of subsidized post-retirement insurance coverage.

We model the retirement decision of 55-64 year-old males using data from two sources: the Current Population Survey (CPS) and the Survey of Income and Program Participation (SIPP). Our major finding, which is robust to a number of specification checks, is a sizeable and significant effect of continuation mandates on retirement decisions: one year of continuation coverage raises retirement rates by 20%. Furthermore, the effect appears to be uniform at all ages rather than disproportionately strong near the age of Medicare eligibility. We also find that continuation mandates had significant effects on the insurance coverage of early retirees. This effect is much larger than the effect on retirement, suggesting that the majority of older individuals with continuation coverage would have retired even in the absence of these benefits.

The organization of the paper is as follows. Section I provides some background on retirement behavior and the health insurance coverage of early retirees. Section II discusses the details of the state and federal continuation of coverage mandates and their likely effects on retirement. The empirical results are presented in Sections III and IV. Section V then focuses on the welfare implications of continuation mandates by considering their impact on insurance coverage. The paper concludes in Section VI with a discussion of the policy implications of our results.

I. Health Insurance and Retirement

A number of studies have examined the retirement behavior of older (primarily male) workers. This literature has focused primarily on the effects of social security (Burtless 1986; Burtless and Moffitt 1984; Diamond and Hausman 1984; Gustman and Steinmeier 1985; Sueyoshi 1989), private pensions (Stock and Wise 1990a and 1990b), and disability insurance (Parsons 1980; Bound 1989). One potentially important factor which has not received much attention is health insurance coverage for retirees. This oversight is especially surprising given the rather consistent evidence that health status is a significant predictor of early retirement (Diamond and Hausman 1984; Bazzoli 1985). If health status matters in the timing of retirement, it seems quite natural that health insurance should matter as well.

Furthermore, individual survey data affirms importance of health insurance in the retirement decision. In a recent Gallup poll, 63 percent of working Americans reported that they "would delay retirement until becoming eligible for Medicare [age 65] if their employers were not going to provide health coverage" despite the fact that 50 percent "said they would prefer to retire early--by age 62" (Employee Benefit Research Institute 1990). This response should not be surprising. 55-64 year-olds are almost three times as likely as individuals aged 25-54 years old to report themselves in fair or poor health; the likelihood of having cancer or a heart attack is four times greater for those aged 55-64 than for those who are 25-54; and the medical expenses of older individuals are almost twice as large, and twice as variable, as those of their younger counterparts. In fact, a one standard deviation increase in expenditures for a 55-64 year old

individual represents 16.5% of average family income for this group; total *family* medical expenditures would naturally constitute a much higher fraction of income.[1]

Despite their higher medical costs, the extent of insurance coverage among 55-64 year-olds is similar to that of 25-54 year olds (Table 1). The sources of health insurance coverage, however, differ with age. Even though employment-based health insurance is the predominant source of coverage regardless of age, older individuals are less likely than younger persons to have this type of insurance and much more likely to be covered by a nongroup (individual) or other group policy. This suggests that individuals who retire early but who do not have access to employer-provided health insurance turn to the individual market for insurance.

The bottom two sections of Table 1 break down the sources of health insurance coverage by employment status. Older individuals who are not employed are 40% less likely to be uninsured than their younger counterparts and six times more likely to be covered by employer-provided health insurance in their own name. These differences are explained in large part by the availability of employer-provided post-retirement health insurance; 45% of individuals work in firms that provide retiree health insurance benefits.[2] There are, nevertheless, a substantial number of older individuals who are not covered by either employer or government-provided

---

[1] Authors' tabulations from the 1987 National Medical Expenditure Survey (health status) and the 1980 National Medical Care and Utilization Expenditure Survey (expenditures). See Gruber and Madrian (1993a) for a more detailed comparison of the health status and medical expenditures of older and younger persons.

[2] See Madrian (1993) for background on the structure and availability of post-retirement health insurance.

health insurance. It is these 30% of non-working older persons who we would expect to benefit from the availability of continuation coverage.[3]

Four recent working papers study the effect of health insurance on retirement. Gustman and Steinmeier (1993) use the Retirement History Survey to ascertain whether individuals have employer-provided retiree health insurance, and data from the 1977 National Medical Care Expenditure Survey to impute the value of that insurance based on individual characteristics. They find that employer-provided health insurance lowers the average age of retirement by 1.3 months. Rogowski and Karoly (1993) use the Survey of Income and Program Participation to impute the provision of retiree health insurance based on firm size and industry and find that retiree coverage raises the retirement hazard by 50%. Lumsdaine, Stock, and Wise (1993) find that Medicare entitlement does not appear to have much explanatory power for the "excess" retirement observed at age 65. Finally, using data from the SIPP and the 1987 National Medical Expenditure Survey, Madrian (1993) shows that individuals with employer provide post-retirement health insurance retire 6-12 months earlier than individuals without such coverage.

One problem that all of these studies face, however, is that they may be unable to control for job characteristics which are correlated with both the generosity of retiree health coverage and the incentives that these jobs offer for retirement. An obvious example is pensions, which are accounted for (in varying degrees) in the first three studies. There may be a number of other ways in which firms encourage or discourage retirement, however, such as through the tasks that they assign older workers or the wage profile that these workers are offered. Furthermore, there may be sorting of workers by retirement propensities into the types of firms that do or do not

---

[3] This figure is a lower bound on the fraction of persons who would benefit from continuation coverage since a number of those reporting employment-based insurance coverage may already be availing themselves of continuation benefits.

offer retiree health insurance. To the extent that these are unobserved to the econometrician but correlated with the offering of retiree coverage, they will bias the estimated effect of such coverage on retirement. What is needed to identify the effect of retiree health insurance is exogenous assignment of such coverage to individuals that is independent of these other job characteristics. Continuation mandates potentially provide such exogenous assignment, and our specification checks below suggest that the assumption of exogeneity is a reasonable one.

II. Continuation of Coverage Laws

For those individuals whose employers do not offer retiree health insurance, an alternative to purchasing health insurance in the individual market is mandated by various state and federal continuation of coverage laws. These laws state that employers sponsoring group health insurance plans must offer terminating employees and their families the right to continue their health insurance coverage through the employer's plan for a specified period of time. The laws generally apply to all separations (except those due to an employees gross misconduct), although in some states benefits are restricted to those who leave their jobs involuntarily.[4] They often also provide benefits to divorced or widowed spouses and their families. The first such law was implemented by Minnesota in 1974. More than 20 states passed similar laws over the next decade before the federal government mandated such coverage at the national level under COBRA.

[4] Because retirement is a voluntary separation, we treat those states whose laws apply only to involuntarily terminated employees as states without laws.

The various state statutes are summarized in Table 2.[5] The length of coverage is generally quite short, from 3-6 months, although 10 states mandate coverage of nine months or more. Most state laws stipulate that an employee must have been covered by an employer's insurance for 3-6 months before being eligible for continuation coverage. This condition, however, is not likely to be binding of older workers, most of whom have been with their current employer for many years.[6] The states laws also apply only to firms that actually purchase insurance through an insurance company; self-insured firms, under the 1974 Employee Retirement Income and Security Act (ERISA), are not subject to these (or any other) state mandates, although they are subject to Federal regulations. The Federal law only applies to firms with more than twenty employees.

Both the state and federal laws stipulate that the employee must pay the full cost of the coverage. At the federal level, this is defined specifically as 102% of the average (not age-specific) employer cost of providing coverage.[7] The coverage must be identical to that provided to similarly situated active employees, including the option to continue enrollment in supplemental insurance plans (such as for vision or dental care) if these are available. Although 102% of the average employer cost is typically much more than individuals pay as active employees, it is substantially less than the cost of buying equivalent coverage in the private market for older workers, for at least two reasons. First, workplace pooling, by reducing

---

[5] Details on state laws are from Hewitt (1985) and Thompson Publishing (1992) and have been cross-checked against the actual statutes. Table 3 lists only those states with laws that apply to employees who terminate their employment voluntarily.

[6] Almost 95% of retirees have job tenure of at least ten years by the time they retire (Madrian 1993).

[7] The employee cost of the state mandates is not made explicit in some state statutes. When it is explicitly legislated, however, the cost sharing rule closely follows the federal formulation.

administrative expenses and the possibilities for adverse selection, is estimated to lower the cost of health insurance in large (10,000 or more employee) firms relative to small (1-4 employee) firms by 40% (Congressional Research Service 1988). Presumably, the cost differentials with individual coverage are even higher. For older individuals, the cost differential between employer-provided and individual health insurance will be further exacerbated by the fact that policies in the individual market are typically age-rated while continuation coverage is not, so that younger and more healthy workers subsidize its cost. The Congressional Research Service (1988) reports that the cost to employers of providing insurance coverage for 55-64 year old males is three times that of providing coverage to males under 40; for females, the ratio is two to one.

In Massachusetts, the average cost of family health insurance coverage per employee in 1993 was $5047.[8] In contrast, a New England commercial insurance company is offering a family policy for a 58 year-old male with a one-year preexisting conditions exclusion for $8640. This represents 26% of the average family income of retired individuals aged 55-64 in Massachusetts. Furthermore, individual policies are typically less generous than employer-provided health insurance; those with group policies are more than 50% more likely to receive ambulance, mental health, and outpatient diagnostic service coverage, and they face lower copayments and deductibles (Farley 1986). Individual policies are also generally medically underwritten so that sick individuals may face substantially higher prices. Alternatively, they may not be able to purchase a policy at all or may face even stricter preexisting conditions exclusions.

[8] Authors' calculation using unpublished data from the Health Insurance Association of America for 1989 (inflated to $1993 using the Medical Care CPI).

The attractiveness of continuation coverage for early retirees is reflected in the high takeup rate of this coverage among those eligible. Flynn (1992) uses data from a large firm that administered COBRA claims over the 1987-1991 period to estimate that 23% of individuals who qualified for COBRA coverage because of retirement elected to receive benefits. While this figure may appear low, it must be interpreted relative to the number of individuals who could be expected to take up such coverage. 20% or early retirees were not insured at their former jobs and are thus not eligible for continuation coverage, while 52% (or 65% of those who had health insurance before they retired) are covered by employer-provided retiree health insurance (Zedlewski 1993). Since continuation benefits are dominated by retiree health insurance in almost all cases, only the 35% of retirees who worked at firms that provided health insurance but who did not receive retiree health insurance should be expected to take up these benefits. This implies that 66% (23%/35%) of those most likely to be covered by continuation benefits actually are. It therefore appears that older workers who retire early and who do not have an alternative source of coverage actually avail themselves of the continuation benefits to which they are entitled.

What effect should we expect continuation benefits to have on the retirement decision?[9] Essentially, continuation benefits provide a subsidy whose value is the certainty equivalent of the difference between the cost of eighteen months of group health insurance and eighteen months of self-insuring, be it through purchasing private health insurance or actually going without. If leisure is a normal good, this subsidy will lead individuals to retire earlier than they would otherwise. The value of this subsidy, and thus the incentive for early retirement, will rise as individuals are sicker or more risk averse.

---

[9] This discussion follows the analytical model developed in Gruber and Madrian (1993a).

The age pattern of these retirement effects is unclear ex ante. If individuals were infinitely risk averse or if coverage in the private market were infinitely expensive, then these laws would necessarily act only as a "bridge to Medicare," allowing individuals to retire a certain number of months before age 65 without exposing themselves to period without group coverage. However, as we showed in Table 1, 30% of early retirees appear willing to face the vagaries of the individual insurance (or self-insurance) market. There are a number of forces, therefore, which will determine the ultimate pattern of age effects.

First, the value of the continuation subsidy will rise with age as the risk of medical expenditures rises. On the other hand, younger retirees may use continuation benefits not as a bridge to Medicare but as a bridge to another job. Diamond and Hausman (1984) report one-year reentry rates for 55-64 year-old retirees of approximately 15%. Sueyoshi (1989) finds that one-third individuals "partially retire," moving from permanent employment to less than full-time work. Higher reentry rates for younger retirees may make continuation benefits more valuable as a means of financing insurance coverage during job changes.

Finally, the analysis has assumed that continuation mandates offer pure rents to workers in the firms that offer this type of coverage. In labor market equilibrium, presumable at least a portion of these rents will be reflected in lower wages for workers with continuation benefits. The extent to which these compensating differentials offset the benefits of retiree health insurance at each age will be a function of the employer's ability to set relative age-specific wages freely,[10] the mobility of workers across firms at different ages, and the excess of the cost of

---

[10] See Rosen (1986) for a discussion of the theory of compensating differentials. Gruber (1993) provides some evidence that shifting the costs of employer-provided benefits to distinct demographic groups in the workplace is feasible.

continuation benefits over the group premium paid by the early retiree.[11] Unlike the pure income effects from continuation benefits, compensating differentials will have both income and substitution effects on the retirement decision leading to uncertain effects on retirement at all ages.

III. Empirical Results from the Current Population Survey

A) Data and Regression Framework

Our primary data source for modelling the transition to retirement is the March Current Population Survey (CPS) for the years 1980-1990. The CPS is a nationally representative survey of over 50,000 households each month, and it has two major advantages for the purpose of this study. First, it is the largest annual data set with information on retirement behavior, which may be important for estimating the effects of state-level laws on a subgroup of workers. Second, it covers the time period of the passage of most state and the federal mandates.[12] Our sample consists of roughly 5000 55-64 year old men from each year, for a total sample size of 56,180.

Each March, the CPS asks not only about current labor force attachment, but about labor force attachment in the previous year as well, including a number of questions about the characteristics of the longest job held during that year. We can therefore observe transitions into retirement over a one year interval. We define individuals as retiring if they worked at least one

---

[11] Huth (1991) reports that the health insurance claims of COBRA recipients exceed those of active employees by 50%. This difference in costs is attributed to adverse selection; it is the sickest individuals who will find continuation coverage most attractive and they will therefore by the ones most likely to take it up. Similar evidence is provided in Long and Marquis (1992).

[12] Individuals are assumed to be subject to a mandate if it is in place by January 1 of the survey year.

week in the previous year but report being retired at the time of the survey.[13] Focusing on individuals who retired during the year enables us to control for a number of characteristics of the individual's previous job, such as industry, occupation, and whether the worker received health insurance or was covered by a pension.

The first column of Table 3 presents some descriptive statistics for our CPS sample. Eight percent are non-white, almost 88% are married, and 70% have employer provided health insurance. The retirement rate is 7.3%, and the average length of continuation coverage is 6.8 months.

We estimate the following probit model of retirement:

$$Pr(Retire_{ijt}) = \Phi(\alpha + \beta_1 \cdot X_{ijt} + \beta_2 \cdot State_j + \beta_3 \cdot Time_t + \beta_4 \cdot Law_{jt}) \qquad (1)$$

where i indexes individuals, j indexes states, and t indexes time. The vector X_{ijt} contains a set of individual demographic and job characteristics: race, marital status, education, and single year age dummies; eight industry and six occupation controls, a control for having health insurance on the job, and a control for being covered by a pension on the job. $State_j$ is a set of state dummies, and $Time_t$ is a set of year dummies. Law_{jt} is the number of months of continuation coverage available in state j at time t. The state fixed effects control for any time invariant characteristics of a state which may be correlated with the state's propensity to pass continuation legislation.[14] The year dummies control for national trends in retirement behavior which may be correlated

---

[13] The estimated coefficients are very similar if we restrict the sample to individuals who worked at least 10 weeks in the previous year. The results are also similar for all persons who report themselves to be out of the labor force, rather than only retired, in the March interview.

[14] We exclude two states from our sample: Hawaii, which has mandated health insurance for all employees, and West Virginia, for which we were unable to definitively date its continuation mandate.

with the passage of these laws. Thus, the effect of the laws is identified in this model by changes in retirement behavior in states which passed the laws (or which were affected by the Federal law), relative to those which did not, during the period after the laws were passed. Further identifying variation comes from differences across states in the number of months of eligibility which these laws allow.

B) Basic Results

The results from this "flow" probit are presented in Table 4. Neither marital status nor education have a significant effect on the probability of retiring although non-whites have significantly lower retirement rates. Workers who receive employer-provided health insurance on their job are also much less likely to retire. To the extent that these workers don't have retiree health insurance, this supports our contention that insurance affects retirement. However, it may also indicate that individuals are less likely to leave "good" jobs.[15] Having a pension significantly raises the probability of retiring, and although the pension coefficient is small relative to the effect of pension incentives estimated with detailed pension information (Stock and Wise 1990a and 1990b), it is similar to previous estimates that also use a pension dummy

[15] The definition of employer-provided health insurance changed in the March 1988 survey. Prior to that time, it measured coverage from the current employer only; it was then redefined to include coverage from a previous employer as well. This means that individuals with retiree coverage would be counted as uninsured before 1988 and insured thereafter. This is not a problem, however, for individuals who worked in the previous year, since the question about insurance coverage in the CPS is whether you had coverage *at any time* during the previous year. Thus, while the number of 55-64 year-old males with employer provided heath insurance rose from 53% on average before 1988 to 62% after, for those who worked in the previous year there was only a 0.05% change (actually a decline) in the rate of insurance coverage.

variable (Samwick 1993).[16] The age coefficients imply that the probability of retiring increases with age, with a spike at age 62 that is generally attributed to the availability of early retirement benefits under Social Security at that age. The probability of retiring at age 64 is approximately 15% higher than at age 55.

The coefficient on *Months of Coverage* is statistically significant and suggests that one year of continuation coverage increases the probability of retiring by 1.5 percentage points or 20% of the baseline retirement probability. This is a sizeable effect for such limited insurance coverage. We will interpret its magnitude more precisely below. First, in order to assess the validity of this finding, we consider a number of specification checks.

C) Specification Checks

Our identification strategy assumes that the continuation mandates provide exogenous variation in the availability of health insurance coverage; that is, that the presence of these laws is otherwise uncorrelated with the retirement decision. We have included state fixed effects in the basic model in order to control for systematic differences between states that did and didn't pass laws. Yet, there remains the possibility that the laws are endogenous responses to *changes* in retirement propensities among the population. That is, if more individuals are retiring, states may respond by mandating benefits that cover individuals after their retirement.

---

[16] One potential weakness of our study is that we don't have better data on pension incentives. This will only affect our results if pension incentives changed systematically in a way which was correlated with the passage of continuation mandates. While we are unable to empirically investigate changes in pension incentives across states and time, there is an insignificant negative relationship between pension coverage and continuation mandates, suggesting that, if anything, changes in pension coverage would bias our results downward.

There is a natural test for this alternative hypothesis, however. Note that only individuals who have health insurance on the job are eligible to receive continuation benefits. Thus, workers without health insurance on the job provide a control group for assessing the effects of continuation mandates. If these mandates are due to (or even simply correlated with) exogenous changes in retirement propensities, then the laws should be correlated with the retirement propensities of workers both with and without health insurance on the job. However, if the laws are *causing* changes in retirement propensities, this should only affect workers with employer-provided health insurance.

The first row of Table 5 reports the results separately for those with and without employer-provided health insurance on their job in the previous year. The coefficient on the number of months of coverage is 40% greater among those with employer-provided health insurance than for the full sample, while it is negative and insignificant for those without such health insurance.[17] Thus, to the extent that uninsured workers in a given state/year provide a valid control group for other factors determining retirement behavior, this suggests that the continuation mandates were having a causal effect.

As noted above, the effect of continuation coverage regulations will be mitigated by factors such as the availability of retiree health insurance or the extent to which firms self-insure (as they are exempt from state mandates). If these factors covary across states with the availability of continuation benefits, then we may have misestimated the effects of continuation

---

[17] The coefficients are not statistically different, however, due to the higher degree of imprecision with which each is estimated in this split sample. If we restrict all of the regressors, except for months of coverage, to have the same coefficients in the two samples, we find a statistically significant effect of mandates on retirement for those with health insurance relative to those without health insurance.

benefits on retirement.[18] Even if that covariance is zero, the fact that these laws should not affect the retirement decisions of some workers implies that our full population estimates of their effects understate the effect of the mandates on the relevant population.

We account for these other factors by constructing a "corrected" months of coverage variable which adjusts the number of months of continuation coverage available by the likelihood that an individual in a given state will be affected by the mandate. The Appendix details the calculation of this correction factor. Briefly, we use unpublished data from the Health Insurance Association of America (HIAA) and the May 1988 Employee Benefits Supplement to the Current Population Survey to estimate, by state, the fraction of workers actually offered employer-provided health insurance, the fraction in firms that offer retiree health insurance, the fraction in self-insured firms, and the fraction in small (<20 employee) firms. If a state law only is in effect, we adjust the months of coverage to reflect the fact that the state law applies only to those offered insurance, and only to those who do not have retiree health insurance and who do not work in a firm that self-insures. If only the federal law is in effect, we adjust the months of coverage by the fraction of individuals who are offered insurance, the fraction who work in small firms, and the fraction in large firms that offer retiree health insurance. When both a state and federal law are in effect, the corrected months of coverage accounts for the fact that the state law will affect insured individuals working in small firms that do not offer retiree health insurance even though the federal law exempts small firms. When we adjust for these factors, the average

[18] Note that this is not captured in the state fixed effects since what is relevant is the interaction between the presence of these factors and the passage of mandates. Our correction factor ignores changes in these factors over time, but Gruber (1992) and Jensen (1993) provide evidence that employer provided insurance coverage and the decision to self-insure are not responsive to the presence of state mandates.

length of continuation coverage in states that offer coverage falls from 14.4 months to 5.24 months.

In the second row of Table 5 we include this corrected months of coverage variable in our transition probits. While the significance of the result does not change, there is a sizeable effect on the magnitude of the coefficient. In the second column, we once again confine the sample to only those with employer-provided health insurance.[19] In this case, the correction factor is increased, as described in Appendix A, to account for the fact that the full population under study has health insurance. Once again, the coefficient is much larger than before. Using these corrected coefficients, a year of continuation coverage now implies a 4.3 percentage point increase the probability of retirement, which is 59% of the baseline retirement probability for 55-64 year-olds.

Finally, our sample selection rule, that individuals be working in the previous year, introduces a "dynamic self-selection" bias into our estimation, as noted by Diamond and Hausman (1984). Imagine that continuation of coverage availability does induce individuals to retire earlier. Then the set of individuals whom we observe actually working in a state after the law has been in place for a number of years will be less likely to retire in response to the law than would the entire population, because those most likely to respond will have already retired. That is, by selecting on those who are still working, we are less likely to find an effect of these provisions since that very group has shown itself less likely to respond. In a bivariate setting, this left-censoring would therefore lead to a downward bias of our estimates of the effect of

[19] The correction factor is not defined for those without employer-provided health insurance because they by definition will not be affected by continuation mandates.

continuation provisions. In a multivariate context, however, the direction of the bias cannot be signed a priori.

With only time invariant covariates in our model, it would be possible to account for this left-censoring by modelling the retirement decisions of those already retired at the start of our sample (Amemiya 1985). Our key covariate, however, the months of continuation coverage available, varies over time. Without knowing its value at the point at which individuals made their retirement decisions, it is not possible to model those decisions.

There is, however, a natural test for the importance of this self-selection bias. Note that the bias should not exist in the first year that a continuation mandate is passed since the sample has not yet been able to select itself based on its response to the mandate. Thus, in the third row of Table 5, we rerun the basic model for only the first year that each state or the federal law is in place. This significantly reduces the number of observations used to estimate our retirement effect. Nevertheless, the coefficient is almost identical to that in the basic specification, suggesting that this left-censoring does not significantly bias our earlier results.[20]

D) Interpretation

To place these findings in context, it is useful to compare them to the estimated increase in retirement propensities following an increase in post-retirement income. In the dynamic stochastic programming model employed by Stock and Wise (1990a and 1990b) and Lumsdaine et al. (1993), a $5000 increase in the value of pension wealth leads to an increase in the

---

[20] Further evidence against the importance of this left-censoring bias is provided in Gruber and Madrian (1993a) which finds that retirement status at a point in time is strongly correlated with these continuation mandates. Furthermore, the estimated effect of continuation coverage on the probability of being retired from the static model in that paper is of almost exactly the same magnitude as the estimated effect on the probability of retiring in Table 5.

retirement hazard of 11.1% for individuals between the ages of 55 and 64.[21] Our basic specification suggests that one year of continuation coverage raises the retirement hazard by 20%, while in our corrected specification the increase in the hazard is as large as 60%. This implies that a year of continuation benefits is valued at between $9,000 and $27,000 in terms of post-retirement wealth. Based on the cost information reported in Section II, a COBRA policy would save an older worker approximately $3,600 per year on the price of family coverage. Taken at face value, our results suggest that workers value the insurance received from continuation policies at a much higher level than its associated cost savings. This may reflect the fact that the individual policy we priced, as with most individual policies, excluded preexisting conditions for some period, thereby understating the true cost difference for an equivalent individual policy. Alternatively, it may be that a number of early retirees must pay substantially more for individual policies or are unable to obtain them at all.

E) Age-Specific Effects

In Section II we discussed the possibility that continuation mandates may have different effects at different ages. We investigate this empirically in Table 6 by allowing the effect of continuation mandates to differ by age. In the first column of Table 6 we see no obvious pattern to the age coefficients, and we cannot reject the restriction that the coefficients are the same at all ages between 55 and 64[22]. The three largest effects occur at ages 58, 60, and 63; the effects at 57, 61, and 64 are also significant at the 10% level.

---

[21] We are grateful to Robin Lumsdaine for performing these calculations for us.

[22] The chi-squared test statistic is 7.65, which is not significant at the 10% level.

In the second and third columns of Table 6, we look separately at the age-specific effects for those with and without health insurance. The results for those with health insurance are very strong at each age and are significant at the 10% level for nine of the ten age interactions. On the other hand, the results for those without health insurance are uniformly insignificant and generally wrong-signed. Once again these results support our contention that the effect of continuation mandates is only operating for those with health insurance on their previous job. As with the full sample, however, there is little pattern evident in the age effects among only those with health insurance. This finding for the pattern of age effects runs counter to the simple presumption that these laws act as a bridge to Medicare. We discuss a possible explanation for this finding below.

V. SIPP Results

A) Data and Model

Several previous studies have suggested that retirement is more naturally modelled in a hazard framework (see for example Diamond and Hausman (1984) and Hausman and Wise (1985)). We have therefore also estimated a hazard model of retirement using data from the 1984-87 Panels of the Survey of Income and Program Participation (SIPP). The SIPP is a longitudinal dataset in which sample members are interviewed every four months for roughly 2½ years and asked to provide information about their labor market activity, income, and participation in welfare and transfer programs over the previous four months. These quarterly interviews are referred to as "Waves". The first interviews of the 1984 Panel (the first SIPP panel available) were conducted in October of 1983, while the initial interviews for subsequent panels commenced in February of the corresponding calendar year. The longitudinal advantage

of the SIPP is, however, countered by two disadvantages; first, the shorter time frame reduces the amount of variation in the state laws, and second, our SIPP sample is less than 8% the size of our CPS sample (roughly 4,400 individuals in the SIPP as opposed to 56,000 in the CPS).[23]

As with the CPS data, we restrict our sample to men aged 55-64. Individuals are included in the sample upon turning 55 and are censored upon reaching their 65th birthday. We exclude individuals who are in the sample for only one wave and individuals already retired when we first observe them. For previously cited reasons, we exclude individuals living in West Virginia and Hawaii. We also drop individuals from several other small states because, out of concern for confidentiality, the SIPP has grouped these states together thereby making it impossible to assign the appropriate state laws to individuals in these states.[24]

Unfortunately, the SIPP does not ask individuals directly whether they have retired. We therefore define retirement to occur when we first see an individual leave the labor force. This has the disadvantage that individuals who are only temporarily out of the labor force will be labelled as retired. We have explored the sensitivity of our findings to alternative definitions such as whether an individual is out of the labor force at two consecutive interview dates (a five month window).[25] Our estimates of the effect of continuation mandates rise substantially when

---

[23] As with other longitudinal data sets, attrition is a potential problem in the SIPP. We treat individuals who leave the panel prematurely as censored observations. As long as the reasons for leaving the sample are not related to the factors that determine retirement, this type of censoring will not bias our results. Klerman (1991) concludes that attrition does not lead to significant bias in measuring health insurance coverage in the SIPP.

[24] These states are Alaska, Idaho, Iowa, Maine, Mississippi, Montana, New Mexico, North Dakota, South Dakota, Vermont and Wyoming. The CPS results are similar if we restrict our CPS sample in the same fashion.

[25] Karoly and Rogowski (1993) use a six month window in the SIPP in their examination of retirement and health insurance.

this alternative rule is imposed. However, the need to censor all individuals at the beginning and end of the sample in order to consistently impose this retirement rule substantially reduces the time frame over which we can observe individuals and further reduces variation in the laws; this removes much of the advantage of turning to the SIPP data.

Figure 1 plots a 5-month moving average of the empirical retirement hazard for our SIPP sample. It looks very similar to retirement hazards based on the Retirement History Survey or the Current Population Survey and shows the characteristic peaks at ages 62 and 65 usually attributed to Social Security eligibility. This gives us some confidence in the reasonableness of our retirement definition.

The means for the SIPP data are presented in the second column of Table 3. Relative to our CPS sample, the SIPP sample is somewhat more likely to be non-white and much more likely to have employer-provided health insurance coverage.[26] The retirement rate in the SIPP is more than double that in the CPS, which is consistent with our observation window in the SIPP being slightly over 2½ years on average.

We use the SIPP to estimate a hazard model of retirement of the type used previously by Diamond and Hausman (1984) and Hausman and Wise (1985). The retirement hazard, $\theta_i(t|\mathbf{x}_i)$, gives the probability that individual i retires in period t conditional on not having already retired:

$$\theta_i(t|\mathbf{x}_i) = \frac{f(t|\mathbf{x}_i)}{1 - F(t|\mathbf{x}_i)} . \qquad (2)$$

We specify a proportional hazards model of retirement,

---

[26] The primary reason for this difference is that a number of individuals without health insurance on their jobs are missing information on the industry and occupation of employment, and are thus excluded from our sample. Our results are similar if these individuals are included in a regression without industry or occupation controls.

$$\theta_i(t|\mathbf{x}_i) = \theta_0(t)\exp\{\mathbf{x}_i'\beta\}, \tag{3}$$

in which age in months, t, is the relevant duration. $\theta_0(t)$, the baseline hazard common to all individuals at time t, is estimated nonparametrically, while the estimates of β are obtained by maximizing the partial likelihood function

$$L(\beta) = \prod_{h=1}^{k} \left[\frac{\exp\{\mathbf{x}_h'\beta\}}{\sum_{l \in R(t)} \exp\{\mathbf{x}_l'\beta\}} \right] \tag{4}$$

where there are k observed exit times and $R(t)$ is the set of all individuals at risk of retiring at time t (Kalbfleisch and Prentice 1980).

As in the CPS, the set of covariates, x, which affect the hazard include marital status, race, education, industry, occupation, and whether or not an individual has health insurance on the job. We include a set of dummies for the year of the SIPP panel in order to capture time trends in retirement, and a dummy for each month in order to model seasonality. We also include a dummy for a whether a particular month is an interview (or "seam") month since there is some tendency in the SIPP for individuals to propagate their current employment status back through the entire quarter covered by the interview. In some specifications we also include dummies for each state; however, the much smaller sample size and shorter time period in the SIPP makes identification from state deviations more tenuous. As with the CPS, the months of continuation coverage is defined as the maximum months of coverage, from the state and federal laws, for which an individual is eligible. This varies by state and may vary over time for the

same individual as state laws begin to take effect or are revised and as the federal law takes effect.[27]

B) Results

Table 7 shows the coefficient estimates and corresponding hazard ratios from our basic hazard specification using data from the SIPP. The hazard ratio gives the effect on the hazard of a one unit change in the corresponding variable; for months of coverage, this is the effect of a one year mandate. Married persons, non-whites, those with more education, and those with health insurance are less likely to retire, but none of the estimated coefficients are significant. The coefficient on months of coverage, however, is statistically significant. It indicates that one year of coverage raises the probability of retirement by 25.2%. This somewhat higher than our CPS estimate. In the second column, which includes state effects, the estimated effect increases to 45.2% with only a moderate increase in the size of the standard error.

In the top row of Table 8, we repeat our earlier specification check of dividing the sample into those who do and do not have health insurance on their job. As before, the effect of continuation coverage is somewhat stronger for those with health insurance and is completely insignificant for those without. Finding this same result from a different dataset and using a different estimation strategy provides additional support for our contention that we are uncovering a causal effect of continuation mandates.[28]

---

[27] We also phase-in the coverage provided under the federal law in 12 equal increments between July 1986 and June 1987 because the mandate at the federal level took effect at the start of a firm's next plan year beginning after July 1986.

[28] "Corrected" estimates in the SIPP, as in the CPS, are equally significant but approximately three times as large.

Finally, in the bottom panel of Table 8 we explore the age pattern of the continuation coverage coefficients in the SIPP. Once again, the presumption that the effects should be strongest near the age of Medicare eligibility is not borne out by the data. In fact, in this case the effects decrease with age, and are even negative at the oldest ages (although insignificant). Note that this decreasing age pattern is not inconsistent with the results in the CPS since these are hazard rates and the CPS findings were probability derivatives; expressed as a hazard, the CPS coefficients yield a similar pattern with age. As with the overall results, the effects are much stronger for those with health insurance than for those without at most ages. Thus, the SIPP confirms the age pattern from the CPS. Continuation mandates are clearly not acting simply as a "bridge to Medicare".[29]

One reason for this, noted in Section II, is that younger "retirees" may be using continuation benefits as a bridge to further employment. If so, our estimates may reflect both the effects of continuation coverage on job turnover and their effects on retirement. In fact, calculations from our SIPP data indicate that of those who leave the labor force for at least one month between the ages of 55 and 62, 40% have returned to the labor force by one year later. Between the ages of 62 and 65, this reentry rate drops to 20%. If younger "retirees" use these benefits as a bridge to other jobs which provide health insurance, they may actually be of more value to that group than to older retirees contemplating a permanent departure from the labor force.[30]

---

[29] It is not clear whether the age pattern is interpreted most naturally as constant (in percentage point terms) or decreasing (in hazard terms).

[30] We are unable to investigate the insurance status on the new job of reentering retirees because the SIPP does not distinguish between insurance from a current employer and a former employer.

This does raise the question of whether our findings are actually effects on retirement, or just effects on turnover. In other research we have found that one year of continuation benefits does increase the likelihood of job turnover among younger workers by about 10% (Gruber and Madrian 1993b). Given that 32% of our SIPP sample reenters on average, this implies that at most 16% of our estimated retirement effect is accounted for by job changers.[31] Furthermore, our definition of retirement in the SIPP is not as stringent as that in the CPS, which requires that individuals actually label themselves retired.[32] The reentry rates estimated from the SIPP, and the bias to our retirement effect from job changes, is therefore most likely an upper bound. Using the 15% reentry rate estimated by Diamond and Hausman (1984), the bias from job changers is only 8%.

VI. Continuation Mandates and Insurance Coverage Among the Retired

If continuation mandates are having an effect on the retirement decisions of older workers, then by definition, they should be affecting their insurance coverage as well. Thus, evidence that such mandates increase insurance coverage among early retirees provides a necessary (but not sufficient) specification check of our result that these mandates affect retirement behavior.

---

[31] A reentry rate of 32% implies that of the 7.3% of individuals in our CPS sample whom we deem to "retire," only 5% will have truly retired while the remaining 2.3% will have left the labor force only temporarily. A 10% increase in the likelihood of changing jobs for the 1.1% who reenter suggests that this group accounts for .23 percentage points (or 16%) of our estimated 1.4 percentage point increase in probability of retiring. Of course, the magnitude of the "turnover bias" to our estimates will be larger for younger workers, although it remains small relative to the retirement effect.

[32] This may explain the larger magnitude of the results in the SIPP; on the other hand, the CPS results are quite similar if we label individuals as retired if they aren't in the labor force for any reason.

Furthermore, the effects of these mandates on insurance coverage may be important for assessing their welfare implications. Ostensibly, the purpose of these mandates was to correct a failure in the private insurance market for job leavers. Such market failure seems likely given the adverse selection in the takeup of continuation coverage benefits documented by Huth (1991) and Long and Marquis (1992). An increase in insurance coverage among individuals who would have retired even in the absence of continuation coverage is an indication that the mandates are having their intended effects in increasing access to group insurance markets. Our finding that one dollar in continuation benefits appears to be valued much more highly than one dollar in post-retirement income suggests that the utility gains from this increased access may be quite large.

On the other hand, the welfare implications of inducing earlier retirement are less clear. To the extent that these mandates cause individuals to retire earlier than they would in a world of no distortions, there is an efficiency cost. This efficiency cost may be small if, as suggested by the results of Kotlikoff and Gohkale (1993), older workers are paid in excess of their marginal product due to implicit long-term employment contracts. But, if the reduction in the wage base necessitates increased taxation of younger workers to maintain a given level of government spending, then even the fall in the wage base (holding productivity constant) may have negative implications. Thus, the welfare effects of the mandates will depend on the relation between their effect on "inframarginal" individuals, who would have retired in their absence, and their effects on the "marginal" individual whose retirement decision is made in response to their presence.

Using the sample of individuals who actually retire in the SIPP, we can model the probability of being covered by employer-provided health insurance after retirement as a function of the months of continuation coverage available at the time of retirement. Our sample consists

of the individuals in our full SIPP sample who retire during our period of observation. The dependent variable is a dummy for whether they have health insurance five months after retirement. The use of this window is dictated by the fact that individuals are only asked about their insurance coverage once per interview, so we must use data from the next interview to assess changes in insurance coverage status. This window is short enough, however, to capture the effects of almost all state laws and the federal law.

We use a simple probit model of the probability that an individual has insurance after retirement. The control variables are the same set that is used in the earlier hazard analysis. As Table 9 shows, the relationship between post-retirement health insurance coverage and months of continuation coverage is significant at the 7% level. The estimate implies that one year of continuation coverage increases the probability of being insured after retirement by 4.7% percentage points. This is 8% of the baseline probability of insurance coverage among early retirees. Compared to our CPS finding that one year of coverage raised retirement rates by 1.4 percentage points, this suggests that the majority of the effect of these mandates is inframarginal; of the individuals gaining coverage from these mandates, the majority would have retired even in their absence.

VI. Conclusions

A number of current public policy proposals in the U.S. would affect the health insurance coverage of early retirees. Thus, it seems especially important at this time to understand the interaction between insurance coverage and the retirement decision. Our strategy for doing so has been to examine the effect of state and federal continuation of coverage mandates on retirement propensities. This has the natural advantage that such coverage was assigned

exogenously to workers in different states at different times, allowing for substantial variation with which to estimate their effects. Our primary finding is that continuation mandates appear to significantly increase retirement propensities. This result emerges from two different data sets using two different estimation strategies and is robust to a variety of specification checks. We also find that these mandates do not simply act as a bridge to Medicare. This may have important implications for assessing the responses of older workers at different ages to future changes in health insurance availability.

Taken at face value, our findings have important implications for public policy. They suggest that increasing the age of full Social Security benefit eligibility to 67 may not have as significant an effect on retirement ages as changing the age of Medicare eligibility. Furthermore, policies to provide health insurance to all citizens in the U.S. could lead to a large increase in the rate of early retirement. This factor should be accounted for in considering the potential financing of such policies. For example, in 1989, 55-64 year old males earned $193 billion.[33] Our results imply that the 18 months of continuation coverage provided under COBRA reduced the 55-64 year old labor force by 2.1% in steady state. If national health insurance has an effect on retirement which is twice as large, then the revenues from financing such health insurance by (for example) a 10% payroll tax would be $800 million smaller than would be apparent from a static revenue calculation.

[33]Authors' tabulations from the March 1990 CPS.

BIBLIOGRAPHY

Amemiya, Takeshi (1985). *Advanced Econometrics*. Cambridge, MA: Harvard University Press.

Bazzoli, Gloria J. (1985). "The Early Retirement Decision: New Empirical Evidence on the Influence of Health." *Journal of Human Resources*, vol. 20, 214-34.

Bound, John (1989). "The Health and Earnings of Rejected Disability Insurance Applicants," *American Economic Review*, vol. 79, 482-503.

Burtless, Gary (1986). "Social Security, Unanticipated Benefit Increases, and the Timing of Retirement." *Review of Economic Studies*, vol. 53, 781-805.

Burtless, Gary and Robert A. Moffitt (1984). "The Effect of Social Security Benefits on the Labor Supply of the Aged." In Henry J. Aaron and Gary Burtless, editors, *Retirement and Economic Behavior*. Washington, DC: Brookings Institution.

Congressional Research Service (1988). *Costs and Effects of Extending Health Insurance Coverage*. Washington, DC: Library of Congress.

Diamond, Peter A. and Jerry A. Hausman (1984). "The Retirement and Unemployment Behavior of Older Men." In Henry J. Aaron and Gary Burtless, editors, *Retirement and Economic Behavior*. Washington, DC: Brookings Institution.

Employee Benefits Research Institute (1990). *Employee Benefit Notes*, November 1990. Washington DC: EBRI.

Farley, Pamela J. (1986). *Private Health Insurance in the United States*, National Health Care Expenditures Study Data Preview 23, DHHS Publication No. (PHS) 86-3406. U.S. Department of Health and Human Services, National Center for Health Services Research and Health Care Technology Assessment.

Flynn, Patrice (1992). "Employment-Based Health Insurance: Coverage Under COBRA Continuation Rules." In U.S. Department of Labor, Pension and Welfare Benefits Administration, *Health Benefits and the Workforce*. Washington, DC: Government Printing Office.

Gruber, Jonathan (1992). "State Mandated Benefits and Employer-Provided Health Insurance." Forthcoming *Journal of Public Economics*.

Gruber, Jonathan (1993). "The Incidence of Mandated Maternity Benefits." Forthcoming *American Economic Review*.

Gruber, Jonathan, and Brigitte Madrian (1993a). "Health Insurance and Early Retirement: Evidence from the Availability of Continuation Coverage." Forthcoming in NBER conference volume on the Economics of Aging, University of Chicago Press.

Gruber, Jonathan, and Brigitte Madrian (1993b). "Limited Insurance Portability and Job Mobility: The Effects of Public Policy on Job-Lock." Massachusetts Institute of Technology mimeo. Cambridge, MA.

Gustman, Alan L. and Thomas L. Steinmeier (1992). "Employer Provided Health Insurance and Retirement Behavior." Dartmouth College mimeo. Hanover, NH.

Hausman, Jerry A. and David A. Wise (1985). "Social Security, Health Status, and Retirement." In David A. Wise, editor, *Pension, Labor and Individual Choice*. Chicago: University of Chicago Press.

Hewitt Associates (1985). *Continuation of Group Medical Coverage--A Study of State Laws*. Lincolnshire, IL: Hewitt Associates.

Hurd, Michael D. (1990). "Research on the Elderly: Economic Status, Retirement, and Consumption and Saving." *Journal of Economic Literature*, vol. 28, 565-637.

Huth, Stephen A. (1991). "COBRA Costs Average 150% of Active Costs." *Employee Benefit Plan Review*, vol. 46, 14-19.

Jensen, Gail (1993). "State Mandated Benefits and the Decision to Self-Insure." Wayne State University mimeo. Detriot, MI.

Kalbfleisch, R.D. and R. L. Prentice (1980). *The Statistical Analysis of Failure Time Data*. New York: Wiley.

Karoly, Lynn A. and Jeannette A. Rogowski, "The Effects of Health Insurance on the Decision to Retire." RAND Corporation mimeo. Santa Monica, CA.

Klerman, Jacob Alex (1991). "Pitfalls of Panel Data: The Case of the SIPP Health Insurance Data." In *Proceedings of the 1991 Public Health Conference on Records and Statistics*. Washington, DC: Public Health Service.

Kotlikoff, Laurence J., and Jagdeesh Gohkale (1992). "Estimating a Firm's Age-Productivity Profile Using the Present Value of Workers' Earnings." *Quarterly Journal of Economics*, vol. 107, 1215-1242.

Long, Stephen H. and M. Susan Marquis (1992). "COBRA Continuation Coverage: Characteristics of Enrollees and Costs in Three Plans." In U.S. Department of Labor, Pension and Welfare Benefits Administration, *Health Benefits and the Workforce*. Washington, DC: Government Printing Office.

Lumsdaine, Robin L., James H. Stock and David A. Wise (1993). "Why Are Retirement Rates So High at Age 65?". Forthcoming in NBER conference volume on the Economics of Aging, University of Chicago Press.

Madrian, Brigitte C. (1993). "Post Retirement Health Insurance and the Decision to Retire." Massachusetts Institute of Technology mimeo. Cambridge, MA.

Parsons, Donald O. (1980). "The Decline of Male Labor Force Participation." *Journal of Political Economy*, vol. 88, 117-134.

Rogowski Jeannette A. and Karoly Lynn A. (1992). "Retirement and Health Insurance Coverage." In U.S. Department of Labor, Pension and Welfare Benefits Administration, *Health Benefits and the Workforce*. Washington, DC: Government Printing Office.

Rosen, Sherwin (1986). "The Theory of Equalizing Differences." In Orley Ashenfelter and Richard Layard, editors, *Handbook of Labor Economics*, Volume I. Amsterdam: North-Holland.

Samwick, Andrew A. (1993). "The Joint Effect of Social Security and Pensions on the Timing of Retirement: Some New Evidence." Massachusetts Institute of Technology mimeo. Cambridge, MA.

Spencer Associates (1991). *Spencer's Research Reports on Employee Benefits*, 329.04.-1 (July 5, 1991). Chicago, IL: Charles D. Spencer & Associates.

Stock, James H. and David A. Wise (1990a). "Pensions, the Option Value of Work, and Retirement." *Econometrica*, vol. 588, 1151-1180.

Stock, James H. and David A. Wise (1990b). "The Pension Inducement to Retire: An Option Value Analysis." In David A. Wise, editor, *Issues in the Economics of Aging*. Chicago: University of Chicago Press.

Sueyoshi, Glenn T. (1989). "Social Security and the Determinants of Full and Partial Retirement: A Competing Risks Analysis." NBER Working Paper No. 3113. Cambridge, MA: National Bureau of Economic Research.

Thompson Publishing Group, Inc. (1992). *Employer's Handbook: Mandated Health Benefits–The COBRA Guide*. Salisbury, MD: Thompson Publishing Group.

Zedlewski, Sheila R. (1993). "Retirees with Employment-Based Health Insurance." In U.S. Department of Labor, Pension and Welfare Benefits Administration, *Trends in Health Benefits*. Washington, DC: Government Printing Office.

APPENDIX A

Calculating Correction Factors for the Impact of State and Federal Continuation Laws

In our basic regression specification, we assign to individuals the maximum number of months of continuation coverage mandated under either federal or state law. There are several factors, however, that lead to less than full coverage of these laws. First, those who are not covered by employer-provided health insurance should not be affected by any of these laws. Similarly, workers in firms that offer retiree health insurance should also be unaffected. Before the federal law took effect, employees of self-insured firms should not have been affected by the state laws because, under the 1974 Employee Retirement Income and Security Act (ERISA), self-insured firms are exempt from state mandates. And last, those who work in firms with fewer than 20 employees will not be affected by the federal law, which does not apply to small firms, although they will be influenced by state laws which pertain to all firms.

To the extent that these factors differ across states, we would expect laws which mandate equivalent months of coverage to have different effects. To account for the less than full coverage of these laws, we compute a "corrected" measure of months of coverage. As before, if there is no state or federal law in effect, we assign no months of coverage to an individual. If a state law is operative but the federal law has not yet taken effect, the corrected months of coverage equals

$$\left[\begin{array}{c}\textit{Months}\\ \textit{of State}\\ \textit{Coverage}\end{array}\right] \cdot \left[\left[\begin{array}{c}\textit{Fraction}\\ \textit{Offered}\\ \textit{Insurance}\end{array}\right] \cdot \left[1 - \begin{array}{c}\textit{Fraction with}\\ \textit{Retiree HI or in}\\ \textit{Self-Insured Firm}\end{array}\right]\right] .$$

If the federal law is in effect but there is no state law, the corrected months of coverage equals

$$\begin{bmatrix} Months \\ of\ Federal \\ Coverage \end{bmatrix} \cdot \left[\begin{bmatrix} Fraction \\ Offered \\ Insurance \end{bmatrix} \cdot \begin{bmatrix} 1 - Fraction\ with \\ Retiree\ HI\ and \\ in > 20\ Firm \end{bmatrix} \cdot \begin{bmatrix} 1 - Fraction \\ Insured\ in \\ < 20\ Firm \end{bmatrix} \right] .$$

Finally, if both a state and the federal law are in place, the corrected months of coverage equals

$$\begin{bmatrix} Months \\ of\ Federal \\ Coverage \end{bmatrix} \cdot \left[\begin{bmatrix} Fraction \\ Offered \\ Insurance \end{bmatrix} \cdot \begin{bmatrix} 1 - Fraction\ with \\ Retiree\ HI\ and \\ in > 20\ Firm \end{bmatrix} \cdot \begin{bmatrix} 1 - Fraction \\ Insured\ in \\ < 20\ Firm \end{bmatrix} \right] +$$

$$\begin{bmatrix} Months \\ of\ State \\ Coverage \end{bmatrix} \cdot \left[\begin{bmatrix} Fraction \\ Offered \\ Insurance \end{bmatrix} \cdot \begin{bmatrix} 1 - Fraction\ with \\ Retiree\ HI\ or\ in \\ Self\text{-}Insured\ Firm \end{bmatrix} \cdot \begin{bmatrix} Fraction \\ Insured\ in \\ < 20\ Firm \end{bmatrix} \right] .$$

We measure the various components of the corrected months of coverage using data from the Health Insurance Association of America (HIAA) 1989 employer survey and the May 1988 Current Population Survey (CPS) pension supplement. In using data from this late date, we are assuming that these factors are constant over time. This is clearly not true for self-insurance which grew dramatically during the 1980s. As long as self-insurance did not grow in a way correlated with the passage of these laws, however, this will not bias our results. Recent research suggests that mandates play little role in the firm's decision to offer insurance or to self-insure; see Gruber (1992) and Jensen (1993).

The fraction of workers offered employer-provided health insurance and the fraction of insured workers in firms with less than 20 employees are measured directly, by state, from the CPS. The fraction with retiree health insurance and in self-insured firms are estimated from the HIAA data. In using the HIAA data, we have tried two different approaches. The first is to calculate the average of these quantities by state. This strategy, however, is problematic, as the state cell sizes in the HIAA data are quite small. Our second approach, therefore, is to use the HIAA data to run a regression predicting the rate of self-insurance and the fraction covered by

retiree health insurance as a function of firm size, industry, and census division. We then use these estimated coefficients to impute values of these quantities for each individual in the May CPS. Our correction factors are state-wide averages of these imputed values.

We did this imputation for two different populations: all male workers, and all male workers over 45 years old. The former yields larger cell sizes in each state, while the latter is closer to the population under study. The correction factors calculated based on these two sample were very highly correlated, so we used only the latter for this analysis.

In our empirical work, we use the corrected months of coverage computed as outlined above when the sample includes all workers. When the sample is restricted to only those with health insurance, we use a similarly constructed corrected months of coverage which does not adjust for the fraction of workers in the state with employer-provided health insurance.

TABLE 1

Insurance Coverage by Age and Employment Status

	Employment-Based		Other	Nongroup	CHAMPUS/	Medicare/	Uninsured
	Any	Own Name	Group		CHAMPVA	Medicaid	
A. All Individuals							
25-54	71.6%	51.1%	1.2%	5.9%	5.7%	5.6	15.4%
55-64	64.5	44.8	4.1	14.5	7.7	10.4	12.0
B. Employed							
25-54	78.5%	62.7%	1.1%	5.8%	4.9%	1.2	13.5%
55-64	76.3	63.1	4.0	12.6	6.8	0.8	10.1
C. Not Employed							
25-54	44.2%	4.2%	1.3%	6.2%	8.8%	23.4	23.0%
55-64	51.6	24.7	4.3	16.6	9.2	20.9	14.1

Source: Authors' calculations using data from the 1987 National Medical Expenditure Survey.

TABLE 2

State Continuation of Coverage Laws

State	Effective Date	Months of Coverage	State	Effective Date	Months of Coverage
Arkansas	7/20/79	4	North Carolina	1/1/82	3
California	1/1/85	3	North Dakota	7/1/83	10
Colorado	7/1/86	3	New York	1/1/86	6
Connecticut	10/1/75	10	Oklahoma	1/1/76	1
	1/1/87	20	Oregon	1/1/82	6
Georgia	7/1/86	3	Rhode Island	1/1/88	18
Illinois	1/1/84	6	South Carolina	1/1/79	2
	8/23/85	9		1/1/90	6
Iowa	7/1/87	9	South Dakota	7/1/84	3
Kansas	1/1/78	6	South Dakota	3/3/88	18
Kentucky	7/15/80	9	Tennessee	1/1/81	3
Minnesota	8/1/74	6	Texas	1/1/81	6
	3/19/83	12	Utah	7/1/86	2
	6/1/87	18	Vermont	5/14/86	6
Missouri	9/28/85	9	Virginia	4/17/86	3
Nevada	1/1/88	18	Wisconsin	5/14/80	18
New Hampshire	8/22/81	10			
New Mexico	7/1/83	6			

Sources: Hewitt (1985), Thompson Publishing Group (1992), and state statutes.

TABLE 3

Descriptive Statistics

	CPS	SIPP
Mean educational attainment	12.0 years	12.3 years
Fraction non-white	7.7%	10.1%
Fraction married	87.8%	88.8%
Fraction with employer-provided health insurance	69.6%	82.9%
Fraction who retire	7.3%	18.6%
Mean months of continuation coverage	6.8 months	7.0 months
Number of individuals	56,180	4,399
Number of monthly observations	-----	77,277

TABLE 4

The Effect of Continuation Coverage on Retirement
Current Population Survey Data

Independent Variable	Coefficient (st. error)	Marginal Prob. $*10^2$
Months of Coverage	.0093	1.497
	(.0047)	
Married	-.0035	-.044
	(.0260)	
Education	.0009	.011
	(.0030)	
Non-white	-.1936	-2.166
	(.0363)	
Health Insurance	-.1220	-1.568
	(.0233)	
Pension	.0621	.772
	(.0220)	
55 Years Old	-1.581	-10.695
	(.1433)	
56 Years Old	-1.522	-10.431
	(.1431)	
57 Years Old	-1.511	-10.296
	(.1431)	
58 Years Old	-1.454	-10.157
	(.1433)	
59 Years Old	-1.338	-9.729
	(.1425)	
60 Years Old	-1.094	-8.751
	(.1419)	
61 Years Old	-.9772	-8.037
	(.1421)	
62 Years Old	-.4638	-4.656
	(.1414)	
63 Years Old	-.4648	-4.600
	(.1414)	
64 Years Old	-.4602	-4.517
	(.1418)	

This table gives estimates from a probit equation for whether or not an individual retired in the last year using data from the 1980-1990 March Current Population Surveys. The sample is comprised of 56,180 men aged 55-64. Coefficients for year, state, industry, and occupation dummies are not reported.

TABLE 5

Coefficient on Months of Coverage Under Alternative Specifications
Current Population Survey Data

Specification	Full Sample		Employer HI Only		No Employer HI Only	
	Coefficient (st. error)	Marginal Prob.*10^2	Coefficient (st. error)	Marginal Prob.*10^2	Coefficient (st. error)	Marginal Prob.*10^2
A. Health Insurance vs. no health insurance	.0093 (.0047)	1.497	.0143 (.0057)	2.225	-.0028 (.0083)	-.414
B. Correct for likelihood of coverage availability	.0240 (.0105)	4.296	.0258 (.0111)	4.684	----	----
C. Test for dynamic self-selection bias	.0089 (.0082)	1.434	.0161 (.0100)	2.547	-.0069 (.0129)	-1.013

This table gives estimates from a probit equation for whether or not an individual retired in the last year using data from the 1980-1990 March Current Population Surveys. The sample is comprised of 56,180 men aged 55-64 of which 39,123 have employer-provided health insurance and 17,057 do not. Coefficients for marital status, education, race, health insurance and pension coverage, age dummies, year, state, industry and occupation dummies are not reported. The second row corrects the months of coverage for state-specific levels of insurance coverage, retiree insurance coverage, self-insurance, and small-firm employment, as described in the Appendix. The third row confines the sample those state/years in which continuation mandates have been in effect for less than a year; there are 34,344 observations in the full sample, 24,177 observations in the health insurance sample, and 10938 observations in the no health insurance sample.

TABLE 6

The Age-Specific Effect of Continuation Coverage on Retirement Transitions
Current Population Survey Data

Independent Variables	Full Sample		Employer HI Only		No Employer HI Only	
	Coefficient (st. error)	Marginal Prob. *10^2	Coefficient (st. error)	Marginal Prob. *10^2	Coefficient (st. error)	Marginal Prob. *10^2
Model 1						
Months of Coverage	.0093 (.0047)	1.497	.0143 (.0057)	2.225	-.0028 (.0083)	-0.414
Model 2						
55*Months	.0061 (.0220)	0.949	.0089 (.0072)	1.337	-.0018 (.0113)	-0.278
56*Months	.0073 (.0060)	1.142	.0132 (.0072)	2.042	-.0087 (.0110)	-1.257
57*Months	.0107 (.0060)	1.731	.0156 (.0072)	2.444	-.0030 (.0109)	-0.455
58*Months	.0116 (.0059)	1.881	.0170 (.0071)	2.694	-.0032 (.0109)	-0.484
59*Months	.0060 (.0057)	0.942	.0138 (.0070)	2.140	-.0122 (.0103)	-1.723
60*Months	.0123 (.0055)	1.977	.0166 (.0067)	2.608	-.0006 (.0101)	-0.095
61*Months	.0097 (.0054)	1.544	.0133 (.0067)	2.040	-.0007 (.0097)	-0.109
62*Months	.0059 (.0052)	0.917	.0107 (.0063)	1.617	-.0050 (.0093)	-0.743
63*Months	.0134 (.0053)	2.198	.0167 (.0065)	2.636	.0051 (.0093)	0.816
64*Months	.0089 (.0054)	1.407	.0181 (.0068)	2.881	-.0048 (.0094)	-0.716

This table gives estimates from a probit equation for whether or not an individual retired in the last year using data from the 1980-1990 March Current Population Surveys. The sample is comprised of 56,180 men aged 55-64 of which 39,123 have employer-provided health insurance and 17,057 do not. Coefficients for marital status, education, race, health insurance and pension coverage, age dummies, year, state, industry and occupation dummies are not reported. The bottom panel gives coefficients from a separate regression which interacts the months of coverage available with a dummy for each separate age.

TABLE 7

The Effect of Continuation Coverage on Retirement
Survey of Income and Program Participation Data

Explanatory Variables	No State Effects		State Effects	
	Coefficient (st. error)	Hazard Ratio	Coefficient (st. error)	Hazard Ratio
Months of coverage	.0187 (.0069)	1.251	.0311 (.0082)	1.452
Married	-.0849 (.1077)	.919	-.1136 (.1084)	.893
Education	-.0104 (.0101)	.990	-.0087 (.0102)	.991
Non-white	-.1385 (.1226)	.871	-.1889 (.1243)	.828
Health insurance	-.0589 (.0961)	.943	-.0620 (.0969)	.940

The table shows estimates of a proportional hazard model of age at retirement for 4,399 men aged 55-64 using data from the 1984, 1985, 1986 and 1987 Panels of the Survey of Income and Program Participation. Coefficients for industry, occupation, month, panel, seam and state dummies are not reported.

TABLE 8

The Age-Specific Effect of Continuation Coverage on Retirement
Survey of Income and Program Participation Data

	Full Sample		Have Health Insurance		No Health Insurance	
	Coefficient (st. error)	Hazard Ratio	Coefficient (st. error)	Hazard Ratio	Coefficient (st. error)	Hazard Ratio
Model 1						
Months of Coverage	.0187 (.0069)	1.251	.0232 (.0075)	1.321	.0033 (.0184)	1.040
Model 2						
55*Months	.0636 (.0159)	2.145	.0688 (.0180)	2.283	.0507 (.0351)	1.837
56*Months	.0481 (.0158)	1.781	.0689 (.0179)	2.286	-.0196 (.0367)	.790
57*Months	.0538 (.0166)	1.907	.0526 (.0180)	1.880	.0575 (.0443)	1.994
58*Months	.0707 (.0143)	2.335	.0716 (.0156)	2.361	.0652 (.0370)	2.187
59*Months	.0256 (.0169)	1.360	.0303 (.0179)	1.438	-.0052 (.0531)	.940
60*Months	.0009 (.0154)	1.011	.0039 (.0166)	1.048	.0004 (.0426)	1.005
61*Months	.0163 (.0132)	1.216	.0210 (.0152)	1.287	.0106 (.0295)	1.136
62*Months	-.0070 (.0119)	.919	-.0040 (.0129)	.953	-.0203 (.0355)	.784
63*Months	-.0015 (.0136)	.982	.0073 (.0144)	1.091	-.0690 (.0491)	.437
64*Months	-.0192 (.0151)	.794	-.0143 (.0169)	.842	-.0188 (.0363)	.798

The table shows estimates of a proportional hazard model of age at retirement for 4,399 men aged 55-64 using data from the 1984, 1985, 1986 and 1987 Panels of the Survey of Income and Program Participation. Coefficients for marital status, education, race, health insurance coverage, industry, occupation, month, panel, and seam dummies are not reported. The bottom panel gives coefficients from a separate regression which interacts the months of coverage available with a dummy for each separate age.

TABLE 9

Continuation Coverage and the Probability of Being Insured After Retirement

Independent Variable	Coefficient (st. error)	Marginal Probability
Married	.1311 (.1714)	.0371
Black	-.4924 (.1822)	-.1487
Education	.0189 (.0159)	.0052
Age	.6705 (.9356)	-.0044
Age^2	-.0056 (.0078)	--
HI on job from which retired	1.742 (.1643)	.6190
Months of coverage	.0139 (.0076)	.0465

The table gives estimates of the probability of being insured after retirement using data from the Survey of Income and Program Participation. The sample is comprised of 937 men aged 55-64 who retire over the sample period. Coefficients for industry, occupation and panel dummies are not reported.

Figure 1: Retirement Hazard in the SIPP